



DEPARTMENT OF ECONOMICS

ISSN 1441-5429

DISCUSSION PAPER 37/11

## Testing for Explosive Behaviour in Relative Inflation Measures: Implications for Monetary Policy<sup>\*</sup>

Vipin Arora, Pedro Gomis-Porqueras<sup>†</sup> and Shuping Shi<sup>‡</sup>

### Abstract

In this paper we test for large deviations in headline measures of the price level relative to core measures using the recently proposed test of [Phillips et al. \(2011a\)](#). We find evidence of explosive behaviour in the headline price index of personal consumption expenditures (PCE) relative to the core PCE (less food and energy prices) on three occasions from 1982-2010. Two of these episodes correspond to energy supply shocks (OPEC price collapse of 1986 and Hurricane Katrina). The third one is during March 2008 through September 2008 which seems to be driven by both food and energy prices as these indices exhibit explosive behaviour. We also find evidence suggesting that inflation expectations behave differently under normal and explosive periods. In particular, unemployment and interest rates also help predict inflation expectations during explosive episodes relative to normal times. Furthermore, explosive episodes in the relative measure between headline and core inflation is found to be more important than the relative volatile periods implied by a Markov-switching model when studying inflation expectations. The findings of this paper suggest that explosive behaviour of headline versus core PCE should be taken into account when conducting monetary policy as it is a key determinant in consumers' inflation expectations.

**Keywords:** Explosive behaviour, core inflation, relative measure, inflation expectations.

**JEL classification:** C5, E31.

---

<sup>\*</sup> We would like to thank Eric Leeper for his comments and suggestions.

<sup>†</sup> Pedro Gomis-Porqueras, Department of Economics, Monash University; Email: [pedro.gomis@monash.edu](mailto:pedro.gomis@monash.edu).

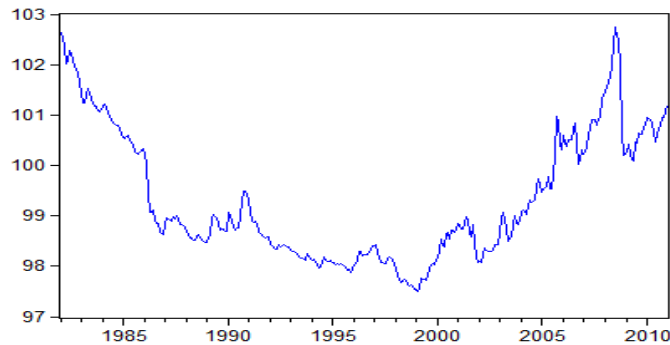
<sup>‡</sup> Shuping Shi, School of Finance, Actuarial Studies and Applied Statistics, The Australian National University; Email: [shuping.shi@anu.edu.au](mailto:shuping.shi@anu.edu.au).

# 1 Introduction

Many central bankers focus on measures of core inflation which exclude certain items that face volatile price movements, notably food and energy. The Federal Reserve, for instance, closely monitors the rate of growth of the core personal consumption expenditure (PCE) deflator.<sup>1</sup> The main argument for central bankers focusing on this narrower measure of inflation is that it can help prevent them from responding too strongly to transitory movements in inflation. Measures of core inflation attempt to strip out or smooth volatile changes in particular prices to distinguish the inflation signal from transitory noise. But how should central bankers react when there are substantial and prolonged differences between core and non-core measures of inflation?

The recent behaviour of commodity prices puts this question into perspective. The prices of food and energy have recently shown substantial volatility and have begun to rise as the global economy recovers. Most other prices do not seem to be rising much, if at all. Figure 1 highlights this fact by plotting the PCE index deflated by the core PCE index. Potential inflationary pressures driven by recent food and energy prices have revived the debate on whether to use a headline or core measure of inflation as Bullard (2011) discusses in his May 2011 speech. This recent and important debate motivates our work.

Figure 1: Headline PCE (deflated by core PCE) from January 1982 to December 2010



One consideration that necessarily enters into this debate is the importance of any differences between the two measures of inflation. As shown in Figure 1, the behaviour of the headline versus core measures can be quite different. Trehan (2011) argues that households are more sensitive to changes in commodity prices and tend to respond by revising their inflation expectations by more than historical relationships warrant. Moreover, relative price movements also convey important information about the scarcity of particular goods and services transmitting vital information necessary for the efficient allocation of resources throughout any market economy. Commodities like oil affect the production and distribution costs of a very wide range of other goods and services, and consequently many other prices. If there are large increases (explosive behaviour) in food and energy prices it is less likely that households will be paying attention to core inflation when forming their inflationary expectations. Unusual patterns for commodity and energy prices do not necessarily affect the ability of central banks to control inflation, but they can greatly complicate the conduct of monetary policy.

Thus, it is plausible that the further food and energy prices are from other prices in the economy, the more difficult it will be for a central bank to anchor inflation expectations. This is the case as relative price movements can have an important impact on the public's inflation expectations (Bullard, 2011).

---

<sup>1</sup>Since February 2000 in the Monetary Policy Report to the Congress, the Federal Reserve Board reports the projections of Federal Open Market Committee participants regarding core PCE inflation, not headline inflation which include food and energy prices.

Consumers confront individual prices, not price indexes, and might interpret big changes in energy or food items as signals of emerging inflation, changing their expectations about future inflation. And when people expect inflation, central banks can find achieving and maintaining price stability more difficult. This possibility poses a policy challenge for central bankers since the impact of monetary policy depends on these inflation expectations remaining “anchored”. When expectations become unanchored, the stance of monetary policy may not be as effective (Bernanke, 2007).

In this paper we provide a procedure that is able to identify the potential differing behaviour between core and headline measures of inflation for the period 1982-2010. In particular, we apply the recently proposed test of Phillips et al. (2011a), PSY hereafter, to identify the periods when headline PCE deviates from core PCE in an explosive manner. Throughout the paper we refer to *explosiveness* as the statistical property of a time series whose characteristic equation has root inside the unit circle. The test of PSY is capable of locating locally (i.e. temporary) explosive behaviour within the sample period. The locating strategy utilises information up to the current period and hence can be used as a warning mechanism for the existence of explosive behaviour. We also isolate the components of the headline measure which are responsible for this behaviour. This is a novel application of the procedure, which generalises the test of Phillips et al. (2011b), PWY hereafter.<sup>2</sup>

Our results indicate that the measure of PCE deviates from core PCE in an explosive manner on three occasions in our sample, two of which may be due to energy supply shocks. From March 2008 to September 2008, both food and energy are behind this rise, as each individual index also shows patterns of explosive movements during this period. Inflation expectations, as measured by surveys, also rise during this period. Additionally, it seems that large energy price movements drive the explosive differences in headline versus core PCE. All three periods of explosive behaviour in the headline series correspond to periods of explosive behaviour in the energy series. This is not true of the food index, as there is a period of explosive behaviour in 2001-2002 that is not represented in the headline measure.

Finally, it is not only the differences between PCE and core PCE measures, but the rate at which one is changing relative to the other that is important. We show using the methods of Mankiw et al. (2003) that expectations are different when headline PCE is deviating explosively from core PCE. Specifically, current and past inflation as well as past unemployment during explosive periods are key in forming inflation expectations relative to normal times. Additionally, we are unable to reject the hypothesis of adaptive expectations during normal periods, but can decisively reject this hypothesis during explosive periods.<sup>3</sup> We also consider a Markov-switching model as an alternative to the previous regime classification when analysing inflation expectations. We find that explosive episodes in the relative measure between headline and core inflation are more important than the relative volatile periods implied by the Markov-switching model when studying inflation expectations. These findings emphasise the need to consider explosive behaviour in the relative measure (which includes the information of both headline and core) when conducting monetary policy.

## 2 Testing for Explosive Deviations

PSY show that the sup augmented Dickey-Fuller (SADF) test of PWY may fail to reveal the existence of explosive behaviour when there are multiple episodes of this behaviour within the same sample period. The generalised sup ADF (GSADF) test of PSY, which was proposed to address this difficulty,

<sup>2</sup>The PSY test was applied to historical stock market data, and identified many periods of explosive behaviour between 1871 and 2010 ((Phillips et al., 2011a)).

<sup>3</sup>In addition, the appendix shows that surveys, model-based forecasts, and market-based measures of inflation expectations all show rises when headline PCE diverges from core PCE in a rapid manner. In particular, survey-based measures respond strongly when headline PCE diverges in an explosive manner from core PCE. This includes both surveys of consumers and surveys of professional forecasters.

significantly improves discriminatory power. In particular, PSY demonstrate via simulations that the GSADF test has significant higher power than the SADF test in identifying the existence of explosive behaviour. Furthermore, they show that when there are multiple explosive episodes in the sample period, the GSADF test can estimate the origination and termination dates of those explosive episodes consistently, whereas the SADF can only consistently estimate dates associated with the first episode.<sup>4</sup>

Before outlining the GSADF test of PSY, we first introduce a backward sup ADF test. The backward sup ADF test implements a right-tailed unit root test (i.e. against an explosive alternative) repeatedly on a backward expanding sample sequence. Suppose  $r_1$  is the (fractional) starting point of a regression sample and  $r_2$  is the (fractional) ending points of the sample. The empirical regression model is

$$\Delta y_t = \alpha_{r_1, r_2} + \beta_{r_1, r_2} y_{t-1} + \sum_{i=1}^k \psi_{r_1, r_2}^i \Delta y_{t-i} + \varepsilon_t, \quad (1)$$

where  $k$  is the lag order and  $\varepsilon_t \stackrel{iid}{\sim} N(0, \sigma_{r_1, r_2}^2)$ . The number of observations in the regression is  $T_w = \lfloor Tr_w \rfloor$ , where  $\lfloor \cdot \rfloor$  signifies the integer part of the argument,  $r_w = r_2 - r_1$  is the (fractional) window size and  $T$  is the sample size. The ADF statistic (t-ratio) based on this regression is denoted by  $ADF_{r_1}^{r_2}$ .

For the backward sup ADF test, the ending point of the samples is fixed at  $r_2$  and the starting point varies from 0 to  $r_2 - r_0$  (it is equivalent to allowing the window size  $r_w$  to expand from  $r_0$  to  $r_2$ ).<sup>5</sup> The backward sup ADF statistic is defined as the sup value of the ADF statistic sequence, denoted by

$$BSADF_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} \{ADF_{r_1}^{r_2}\}.$$

The GSADF test can be viewed as a repeat implementation of a backward sup ADF test for each  $r_2 \in [r_0, 1]$ . The GSADF statistic is defined as the sup value of the backward sup ADF statistic sequence and denoted by<sup>6</sup>

$$GSADF(r_0) = \sup_{r_2 \in [r_0, 1]} \{BSADF_{r_2}(r_0)\}.$$

The GSADF statistic is utilised to conduct inference of the existence of explosive behaviour within the whole sample period. Suppose there is evidence of explosive behaviour, one can then date stamp the occurrence periods using the backward sup ADF statistic. Specifically, we conclude that observation  $\lfloor Tr_2 \rfloor$  belongs to an explosive phase in the trajectory given that

$$BSADF_{r_2}(r_0) > scv^\alpha(r_0),$$

where  $scv^\alpha(r_0)$  is the  $100(1 - \alpha)\%$  (right-tail) critical value of the backward sup ADF statistic.

Notice that the backward sup ADF statistic  $BSADF_{r_2}(r_0)$  is calculated using information up to period  $\lfloor Tr_2 \rfloor$ . It does not depend on future realisations and hence this strategy can serve as a warning

<sup>4</sup>It cannot consistently estimate the origination and termination dates associated with the subsequent episodes.

<sup>5</sup>The minimum window size  $r_0$  is selected to ensure that there are sufficient observations to achieve estimation efficiency.

<sup>6</sup>Under the null hypothesis that  $y_t$  is a random walk with an asymptotically negligible drift (namely  $y_t = dT^{-\eta} + y_{t-1} + \varepsilon_t$ ,  $\varepsilon_t \stackrel{iid}{\sim} N(0, \sigma^2)$ , constant  $d$  and  $\eta > 1/2$ ), the asymptotic distribution of the GSADF statistic is

$$\sup_{r_2 \in [r_0, 1], r_1 \in [0, r_2 - r_0]} \left\{ \frac{\frac{1}{2} r_w [W(r_2)^2 - W(r_1)^2 - r_w] - \int_{r_1}^{r_2} W(r) dr [W(r_2) - W(r_1)]}{r_w^{1/2} \left\{ r_w \int_{r_1}^{r_2} W(r)^2 dr - \left[ \int_{r_1}^{r_2} W(r) dr \right]^2 \right\}^{1/2}} \right\},$$

where  $W$  is the standard Wiener process.

mechanism for explosive behaviour. A more detailed illustration of the GSADF test can be found in Phillips et al. (2011a).<sup>7</sup>

## 2.1 Data

The data series are taken from the Federal Reserve Bank of St. Louis’s Federal Reserve Economic Data (FRED), and the U.S. Bureau of Economic Analysis (BEA) for the period 1982-2010. The PCE headline measure (PCEPI) is seasonally adjusted and at a monthly frequency, and comes from the GDP and components section of FRED. The PCE core measure (PCEPILFE) is also seasonally adjusted and obtained monthly from the same section of FRED, and excludes food and energy. The individual measures of food and energy are taken directly from the BEA, but at a quarterly frequency, and not seasonally adjusted. The food component is “food and beverages purchased for off-premises consumption,” and the energy component is “energy goods and services”.

The headline PCE numbers along with the food and energy indices are all deflated by core PCE, which we refer as relative measure and relative food and energy indices. This gives a measure of each of how each of the respective price indexes are moving relative to core over time.<sup>8</sup> The base year is 2005. In the case of the food and energy indexes a measure of core PCE at a quarterly frequency is used for deflation. The logarithm of each deflated series is then used when performing the tests.

## 2.2 Results

Table 1 shows the generalised SADF statistic, along with respective finite sample critical values.<sup>9</sup> As we can see from Table 1, we find evidence for explosive behaviour in each of the data series tested. In particular, for the relative measure between headline and core PCE, the generalised sup ADF statistic is 3.257, which is greater than the 99% critical value 2.687. This implies the existence of explosive behaviour in the sample period. Similar results hold for the food and energy series as well. Each of generalised SADF statistics are above the 99% critical values. The exact periods of explosive behaviour for each index (relative to core PCE) are depicted in Figures 2, 3, and 4.

Table 1: The generalized sup ADF test

	Relative PCE	Food & Beverage	Energy & Service
GSADF	3.257	3.520	3.592
90%	1.907	2.137	2.137
95%	2.162	2.472	2.472
99%	2.687	3.419	3.419

Note: The critical values are obtained from Monte Carlo simulations with 2,000 replications.

Figure 2 plots the backward SADF statistic sequence against its 90% critical value sequence for the relative PCE. We find explosive behaviour whenever the BSADF statistic exceeds the critical value. The relative measure between headline and core PCE has three such events, FEB86-JAN87, SEP05-OCT05, and MAR08-SEP08. The first two periods of explosive behaviour are likely to be

<sup>7</sup>The Gauss and Matlab programs for implementing this test are available for download from <https://sites.google.com/site/shupingshi/PrgGSADF.zip?attredirects=0&d=1>.

<sup>8</sup>This is the same as taking the ratio of each price index to core PCE each period and multiplying by a constant.

<sup>9</sup>The test results of the generalised SADF test are not sensitive to the lag order in the regression model. In what follows the lag order is set to zero. The minimum window size is 36 for the (monthly) headline PCE index and 12 for the (quarterly) food and energy indexes (i.e. 3 years). The critical values are obtained from Monte Carlo simulations with 2,000 replications (parameters  $d$  and  $\eta$  in the null model are set to unity).

Figure 2: Explosive periods in headline PCE (deflated by core PCE)

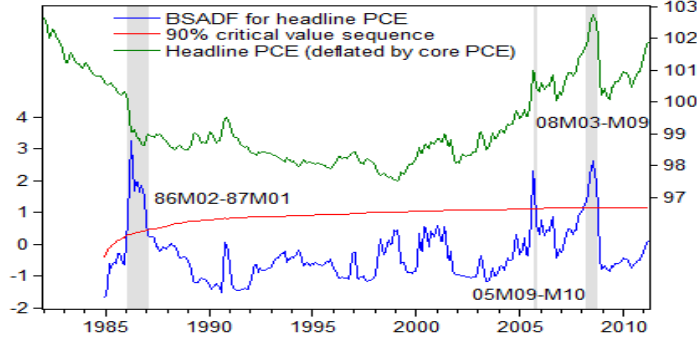
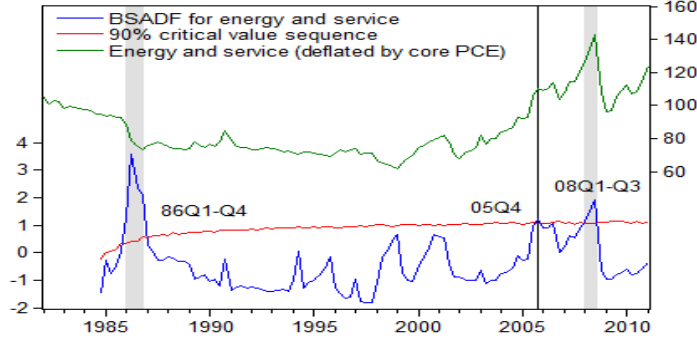
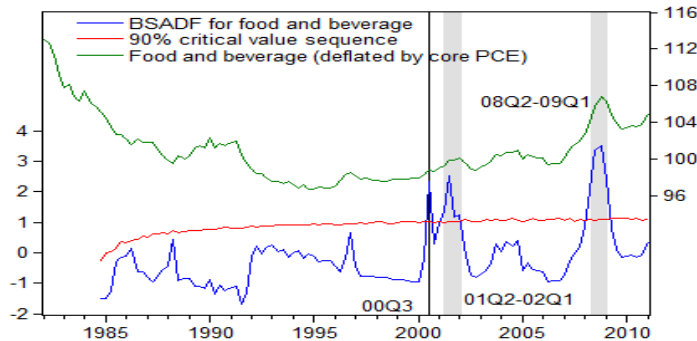


Figure 3: Explosive periods in energy and service components (deflated by core PCE)



driven by energy supply shocks. In particular, the first period may correspond to the OPEC collapse of 1986, which resulted in a surplus of oil on world markets. The second period may reflect the onset of Hurricane Katrina, which shut down many refineries in the Southern United States. The third period is the most dramatic as the explosive behaviour in the relative measure lasts more than half a year.

Figure 4: Explosive periods in food and beverage components (deflated by core PCE)



To gain deeper insight we consider specific time series. Figures 3 and 4 show that both energy and food may seem to be causing this rise. Each of these series also show explosive behaviour in 2008. Figure 3 confirms that periods of explosive behaviour in the relative measure between headline and core PCE were due to energy shocks, as there are rises in energy in 1986 and 2005. Interestingly,

Figure 4 reveals that the explosive behaviour in food in 2000 and 2001-2002 does not translate into similar movements in headline PCE. This suggests that explosive energy price movements are more important than similar movements in food prices in generating explosive behaviour for the aggregate measure of inflation.

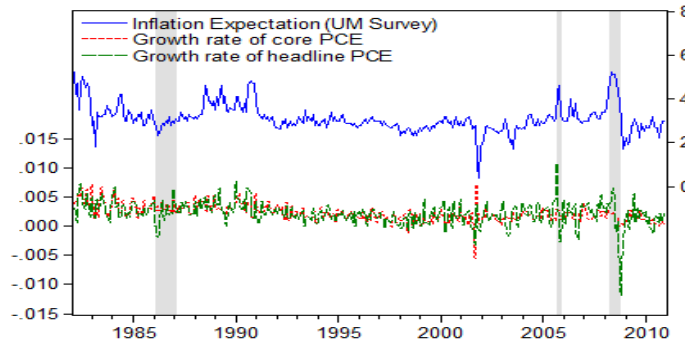
### 3 Explosiveness and Consumer's Inflation Expectations

A straight forward method to measure inflation expectations of consumers is to ask them to present quantitative estimates. For instance, each month, the University of Michigan's Survey Research Center assesses consumer sentiment by interviewing a random sample of approximately 500 U.S. households. As part of the survey, respondents are asked to forecast key macroeconomic variables, such as inflation, interest rates, and unemployment. An alternative is to use prices of index-linked financial securities to provide market-based measures of inflation expectations and attitudes towards inflation risk.

Various types of each measure are available.<sup>10</sup> In this section we focus on estimates from the University of Michigan's Survey of Consumer Attitudes and Behaviour (UM) and explore how these behave in the presence of large deviations in the headline and core inflation measures.

Given the explosive deviations in the relative measure between headline and core PCE shown in the previous section, we next examine some implications for inflation expectations. Here, we argue that these explosive movements are in fact different than other movements in the relative measure between headline and core PCE with regards to their impact on inflation expectations. Figure 5 plots the one-year ahead inflation expectations from the UM survey and the growth rates of headline PCE and core PCE over the period of January 1982 to December 2010. The shaded areas identify the explosive periods in the relative measure found in the previous section.

Figure 5: Inflation expectations and growth rates of headline PCE and core PCE



As we can see, there are sharp and pronounced rises in inflation expectations when relative measure between headline and core PCE deviates explosively *upward* from core PCE (2005 and 2008). In the sample period the only other similar rises in inflation expectations occur around 1990, but the magnitude in this case is smaller. It seems that explosive upward movements in prices may have a different impact on inflation expectations than other periods.

The largest fall in inflation expectations follows the explosive rise in food and beverage prices in 2001-2002. This may be because food and beverages are commonly purchased items, and a fall in

<sup>10</sup>Other estimates include the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters (SPF), the Federal Reserve Bank of Cleveland's monthly model-based inflation expectation measure (FRC), inflation swap rates, and the difference between yields on nominal U.S. Treasury Notes and Treasury Inflation Protected Securities (TIPS) (Pasaogullari, 2011).



these prices have a larger impact on consumer expectations. Moreover, we find that the explosive *downward* deviation of headline from core PCE follows the OPEC collapse of 1985-1986 but does not result in a sharp and pronounced fall in expectations. This asymmetric impact of price changes is line with the findings of Engemann et al. (2011). These authors find that sharp increases in oil prices affect economic activity adversely, but sharp decreases in oil prices have no effect.

Table 2 presents the correlation coefficients between the change of inflation expectations (UM survey) and different observed inflation measures. In particular, the correlation between the change of inflation expectations and the change of the relative inflation measure is 0.432, which is larger than those between inflation expectations and the growth rates of headline PCE and core PCE. Furthermore, the growth rate of headline PCE is more related to inflation expectations than the growth rate of core PCE as it has a larger coefficient. This then seems to suggest that inflation expectations are more closely related to the relative measure of inflation than core inflation itself.

Table 2: Correlations between inflation expectations and inflation measures

	$\Delta$ Relative In- flation	Growth Rate of Headline PCE	Growth Rate of Core PCE
$\Delta$ Inflation Expectations	0.432	0.308	-0.054

Note:  $\Delta$  is the first order difference operator.

While the previous analysis gives some interesting insights into inflation expectations and explosive behaviour in headline PCE, a more formal analysis is needed clarify this relationship. In an influential paper, Mankiw et al. (2003), MRJ hereafter, document some important features of survey-based measures of inflation expectations. In particular, irrespective of the survey used, neither rationally generated nor adaptively generated expectations can fully account for the forecasts of inflation expectations. This latter claim is based on a series of regressions where inflation expectations from different surveys are regressed on possible explanatory variables such as inflation ( $\pi_t$ ), unemployment ( $U_t$ ), and interest rates ( $i_t$ ). More specifically, MRJ propose the following specification

$$E_t\pi_{t+12} = \alpha + \beta(L)\pi_t + \gamma U_t + \kappa U_{t-3} + \delta i_t + \phi i_{t-3} + \varepsilon_t, \quad (2)$$

where  $E_t\pi_{t+12}$  is the twelve-period ahead inflation expectation,  $L$  is lag operator and  $\varepsilon_t$  is the error term. The values of the regression coefficients indicate that survey respondents neither fully incorporate all of the past information ( $\beta(1) < 1$ ), nor do they only use past information on inflation (i.e. MRJ reject the null hypothesis that  $\gamma = \kappa = \delta = \phi = 0$ ).

In this paper, we use the MRJ basic setup to gauge how inflation expectations change during periods of upward explosive deviations in headline PCE. Specifically, we add a dummy variable ( $D_t$ ) that captures periods of upward explosive deviation which we identified in the previous section. In particular,  $D_t$  has a value of 1 during periods of upward explosive deviation of headline PCE from core PCE, and 0 otherwise. This is generated based on the following criteria:

$$D_t = \begin{cases} 1 & \text{if } BSADF_t^h(r_0) \geq scv_t^\alpha(r_0) \text{ and } H_t \geq H_{t-1} \\ 0 & \text{Otherwise} \end{cases},$$

where  $BSADF_t^h(r_0)$  is the BSADF statistic for the relative measure between headline and core PCE,  $scv_t^\alpha(r_0)$  is the finite sample critical value of the statistic and  $H_t$  is the current period relative measure between headline and core PCE.

The proposed regression model to study the relationship between inflation expectations and ex-



plosive inflation behaviour is given by

$$E_t \pi_{t+12} = \alpha + \beta(L) \pi_t + \gamma U_t + \kappa U_{t-1} + \delta i_t + \phi i_{t-1} + \alpha' D_t + \beta(L)' D_t \pi_t + \gamma' D_t U_t + \kappa' D_t U_{t-1} + \delta' D_t i_t + \phi' D_t i_{t-1} + \varepsilon_t. \quad (3)$$

Our regression model includes  $U_{t-1}$  and  $i_{t-1}$  instead of  $U_{t-3}$  and  $i_{t-3}$ , as in MRJ. We do so as the U.S. unemployment and interest rate data are released monthly, providing more data points for our analysis.<sup>11</sup> We include the year-on-year observed inflation for each of the previous three months as well.<sup>12</sup> If expectations change during explosive periods, the coefficients on the dummies should be jointly significant (i.e. we should reject the null hypothesis that  $\alpha' = \beta(1)' = \gamma' = \kappa' = \delta' = \phi' = 0$ ). Following MRJ, we also test for adaptive expectation for the normal periods ( $\gamma = \kappa = \delta = \phi = 0$ ) and for the explosive periods ( $\gamma' = \kappa' = \delta' = \phi' = 0$ ) respectively.

Table 3 reports the estimation and hypothesis test results of a baseline model based on MRJ and our proposed model. The baseline model uses one month lags of the unemployment and interest rates as independent variables to make it consistent with our modification.

Table 3: Test of adaptive expectations for the Michigan median inflation expectations

	Baseline Model		Proposed Model	
Inflation				
$\beta(1)$ : sum of 3 coefficients	0.298***	(0.077)	0.240***	(0.056)
Unemployment				
$\gamma$ : date of forecast	-0.031	(0.057)	-0.048	(0.055)
$\kappa$ : 1 months prior	0.056	(0.060)	0.091*	(0.054)
Treasury bill rate				
$\delta$ : date of forecast	0.259	(0.156)	0.252	(0.161)
$\phi$ : 1 months prior	-0.258	(0.152)	-0.214	(0.151)
Inflation x Dummy				
$\beta(1)'$ : sum of 3 coefficients			0.304***	(0.056)
Unemployment x Dummy				
$\gamma'$ : date of forecast			0.309***	(0.055)
$\kappa'$ : 1 months prior			-0.531***	(0.054)
Treasury bill rate x Dummy				
$\delta'$ : date of forecast			0.272*	(0.161)
$\phi'$ : 1 months prior			-0.408***	(0.151)
Joint significance test of parameters related to dummy			$F_{8,297} = 99.24^{***}$	
Reject adaptive expectation?	$F_{4,302} = 0.892^{(a)}$		$F_{4,297} = 1.39^{(a)}$	
			$F_{4,297}^e = 47.54^{***(b)}$	
Adjusted $R^2$	0.442		0.565	

Note: Parameters in parentheses are the Newey-West standard errors (lag truncation=5). \*\*\*, \*\* and \* denote statistical significance at the 1%, 5% and 10% levels respectively. (a) are F-statistics for testing  $\gamma = \kappa = \delta = \phi = 0$ . (b) is the F-statistic for testing  $\gamma' = \kappa' = \delta' = \phi' = 0$ . The regression starts from December 1984 since the minimum window size in the explosive test is 36.

As we can see from Table 3, we cannot reject the null hypothesis of adaptive expectations over

<sup>11</sup>It is also likely that consumers form their inflation expectations based on the newly released (monthly) data.

<sup>12</sup>This is restricted by the number of observations in the explosive rising regime.

the sample period from the baseline model.<sup>13</sup> Moreover, the dummy variables are jointly (and highly) significant. This indicates that inflation expectations during periods when headline PCE deviates from core PCE upward in an explosive manner are quite relevant when forming expectations. Interestingly, the results indicate that consumers rely more on the past inflation during these explosive phases.

Finally, consistent with the baseline model, we cannot reject adaptive expectations during normal periods. This is in sharp contrast to explosive periods where adaptive expectations are strongly rejected. This difference provides some evidence that expectations change during such periods. Moreover, unemployment and interest rates also help predict inflation expectations during these periods.

## 4 Markov-switching and Consumer's Inflation Expectations

The strategy of PSY used in the previous section implicitly classifies the relative measure between headline and core PCE into an explosive regime and a non-explosive regime. We demonstrate that when there is explosive behaviour in the relative measure between headline and core inflation, consumers depart from adaptive inflation expectation and utilise information other than past inflation to forecast the twelve-period ahead inflation.

In this section, we consider an alternative to the previous regime classification. The new framework assumes that the relative measure has two regimes and these two regimes switch from one to the other in a Markov pattern. More structurally, the Markov-switching model is specified as follows:

$$\Delta y_t = \alpha_{s_t} + \beta_{s_t} y_{t-1} + \sum_{j=1}^2 \psi_{s_t,j} \Delta y_{t-j} + \varepsilon_t, \varepsilon_t \sim N(0, \sigma_{s_t}), \quad (4)$$

where  $y_t$  is the relative measure between headline and core PCE and  $s_t$  is a realisation of a state variable  $S_t$ , takes value 0 or 1. The state variable  $S_t$  is governed by a first order Markov-chain, namely

$$P\{s_t = 0 | s_{t-1} = 0\} = p \text{ and } P\{s_t = 1 | s_{t-1} = 1\} = q.$$

The corresponding log likelihood function<sup>14</sup> is

$$l(y_1, \dots, y_T; \Psi) = \sum_{t=1}^T \log \sum_{s_t \in \{0,1\}} f(y_t | \Upsilon_{t-1}, S_t = s_t; \Psi) \times \Pr\{S_t = s_t | \Upsilon_{t-1}; \Psi\},$$

where  $\Psi$  contains all of the unknown parameters and  $\Upsilon_{t-1}$  is the information set available at period  $t - 1$ . The model is estimated using the Broyden–Fletcher–Goldfarb–Shanno (BFGS) algorithm with 100 sets of randomly generated start-up values, and we choose the one associated with the largest likelihood value. The smoothed probabilities are calculated according to [Kim \(1994\)](#).

Table 4 presents estimates associated with the Markov-switching model. Notice that estimates  $\beta_0$  and  $\beta_1$  are both smaller than zero. Regime 1 has larger conditional mean ( $\alpha_1$ ) and conditional standard deviation ( $\sigma_1$ ) than those of regime 0. In particular,  $\sigma_1$  is 3.23 times larger than  $\sigma_0$ . Furthermore, the likelihood ratio test indicates that  $\sigma_0$  is significantly different from  $\sigma_1$ . The smoothed probabilities of being in regime 1 are displayed in Figure 6. As evident by this figure, there is a high probability of being in regime 1 (greater than 0.5) during the periods of 1986 OPEC collapse, 1990, 2000-2001, and 2003 onwards.

In order to explore how inflation expectations change over these two regimes, we replace the dummy variable  $D_t$  in equation (3) by the smoothed probabilities of being in regime 1 ( $p_t^s$ ). The new model

<sup>13</sup>The estimation and test results of the baseline model are not sensitive to the lag selection of inflation.

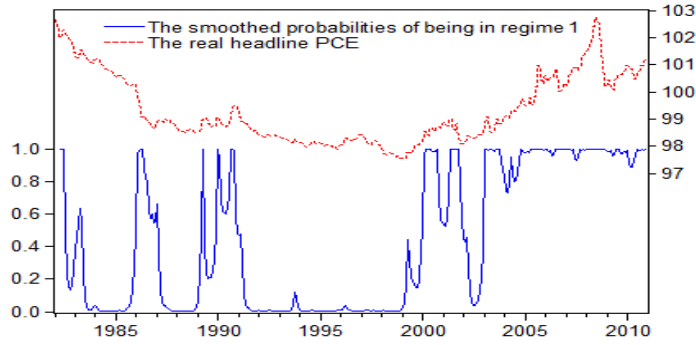
<sup>14</sup>Analogous to [Shi \(2010\)](#), we resort to the Quasi-Bayesian approach ([Hamilton, 1991](#)) to address of the problem of unbounded likelihood ([Day, 1969](#)). However, this adjustment does not change the estimation results of the relative PCE.

Table 4: Estimates of the Markov-switching model: relative measure between headline and core PCE

Regime 0			Regime 1		
$\alpha_0$	1.450	(3.21)	$\alpha_1$	2.526	(1.79)
$\beta_0$	-0.015	(-3.24)	$\beta_1$	-0.025	(-1.78)
$\psi_{01}$	0.368	(3.37)	$\psi_{11}$	0.511	(6.36)
$\psi_{02}$	-0.147	(-1.85)	$\psi_{11}$	-0.262	(-3.25)
$\sigma_0$	0.061	(10.47)	$\sigma_1$	0.197	(14.04)
$p$	0.952	(36.77)	$q$	0.946	(28.42)
Likelihood ratio stat. ( $\sigma_0 = \sigma_1$ )			55.408 [0.000]		

Note: Figures in parentheses are t-statistics. Figures in the square bracket is p-value.

Figure 6: The smoothed probabilities of being in regime 1



for inflation expectation is

$$\begin{aligned}
 E_t \pi_{t+12} = & \alpha + \beta(L) \pi_t + \gamma U_t + \kappa U_{t-1} + \delta i_t + \phi i_{t-1} \\
 & + \alpha' p_t^s + \beta(L)' p_t^s \pi_t + \gamma' p_t^s U_t + \kappa' p_t^s U_{t-1} + \delta' p_t^s i_t + \phi' p_t^s i_{t-1} + \varepsilon_t.
 \end{aligned} \tag{5}$$

If parameters related to the auxiliary variable  $p_t^s$  are jointly significant, it suggests that inflation expectations in regime 0 and regime 1 are different. This could potentially be due to the fact that consumers rely on different information in these two regimes (i.e. use unemployment and interest rate to help predict or not). Alternatively, consumers use the same set of information but the dependence level of their forecast on the information varies across regimes.

The estimation results of model (5), along with related hypothesis tests, are presented in Table 5. We fail to reject the joint insignificance of  $\gamma, \kappa, \delta, \phi, \gamma', \kappa', \delta', \phi'$ . This suggests that consumers do not rely on unemployment and interest rate information to predict inflation in both regimes. Therefore, we cannot reject adaptive expectation in both regimes.<sup>15</sup> This finding is consistent with the baseline model where we cannot reject adaptive expectation for the whole sample period. Nevertheless, it is in sharp contrast to the finding of model (3). According to model (3), adaptive expectations do not hold in periods of explosive divergence of headline PCE from core PCE.

Finally, we note that the adjusted R-square of model (3) is higher than that of model (5). This suggests that the explosiveness indicator (i.e. the dummy variable  $D_t$  obtained from the PSY strategy) performs better than the regime-switching indicator (i.e. the smoothed probability  $p_t^s$  obtained from

<sup>15</sup>Notice that all parameters related to the auxiliary variable are individually insignificant. The joint significance of parameters related to the auxiliary variable indicates that the extend of consumers relying on past inflation changed slightly across these two regimes.

Table 5: Test of adaptive expectations for the Michigan median inflation expectations (cont.)

	Proposed Model (Dummy)		Proposed Model (Smoothed Prob.)	
Inflation				
$\beta(1)$ : sum of 3 coefficients	0.240***	(0.06)	0.282***	(4.48)
Unemployment				
$\gamma$ : date of forecast	-0.048	(0.06)	-0.066	(-1.40)
$\kappa$ : 1 months prior	0.091*	(0.05)	0.064	(1.44)
Treasury bill rate				
$\delta$ : date of forecast	0.252	(0.16)	0.032	(0.23)
$\phi$ : 1 months prior	-0.214	(0.15)	0.025	(0.17)
Inflation x Auxiliary				
$\beta(1)'$ : sum of 3 coefficients	0.304***	(0.06)	0.019	(0.13)
Unemployment x Auxiliary				
$\gamma'$ : date of forecast	0.309***	(0.06)	0.027	(0.19)
$\kappa'$ : 1 months prior	-0.531***	(0.05)	0.012	(0.08)
Treasury bill rate x Auxiliary				
$\delta'$ : date of forecast	0.272*	(0.16)	0.138	(0.42)
$\phi'$ : 1 months prior	-0.408***	(0.15)	-0.195	(-0.57)
Joint significance test of parameters related to auxiliary variable	$F_{8,297} = 99.24^{***}$		$F_{8,329} = 1.99^{**}$	
Reject adaptive expectation?	$F_{4,297} = 1.39^{(a)}$ $F_{4,297}^e = 47.54^{***(b)}$		$F_{8,329} = 1.16^{(c)}$	
Adjusted $R^2$	0.565		0.491	

Note: Parameters in parentheses are the Newey-West standard errors (lag truncation=5). \*\*\*, \*\* and \* denote statistical significance at the 1%, 5% and 10% levels respectively. (a) is the F-statistic for testing  $\gamma = \kappa = \delta = \phi = 0$ . (b) is the F-statistic for testing  $\gamma' = \kappa' = \delta' = \phi' = 0$ . (c) is the F-statistic for testing for the joint significance of  $\gamma, \kappa, \delta, \phi, \gamma', \kappa', \delta'$  and  $\phi'$ .

the Markov-switching model) in capturing the dynamics of consumers' inflation expectations. These findings suggest that inflation expectations can be better forecasted when exploiting the explosive behaviour of the relative measure between headline and core PCE highlighting the potential benefits for monetary policy.

## 5 Conclusion

We apply the recently proposed test of Phillips et al. (2011a) to identify episodes where headline PCE deviates from core PCE in an explosive manner.<sup>16</sup> We also isolate the components of the headline measure which are responsible for this behaviour. We find there has been explosive behaviour in headline PCE relative to core PCE on three occasions since 1982. It also seems that explosive behaviour in the headline PCE series on these occasions is driven by similar behaviour in energy prices. Two of these periods correspond to energy supply shocks (the OPEC collapse of 1986 and Hurricane Katrina). The third period of explosive behaviour was from March 2008 to September 2008.

Finally, we find evidence suggesting that inflation expectations behave differently under normal

<sup>16</sup>Other test focusing on specific commodity prices have been done by Shi and Arora (2011), Phillips and Yu (2011), and Gilbert (2009).

and explosive periods. In particular, unemployment and interest rates also help predict inflation expectations during explosive episodes relative to normal times. Additionally, we consider a Markov-switching process for the relative measure between headline and core inflation when analyzing inflation expectations. We find that explosive deviations of the relative measure are more important than the relative volatility implied by the Markov switching model when studying inflation expectations. The findings of this paper suggest that explosive behaviour of headline PCE deflated by core PCE should be taken into account when conducting monetary policy as it is a key determinant in consumers' inflation expectations.

## References

- Bernanke, Ben S.**, “Inflation Expectations and Inflation Forecasting,” <http://www.federalreserve.gov/newsevents/speech/bernanke20070710a.htm> 2007. At the Monetary Economics Workshop of the National Bureau of Economic Research Summer Institute, Cambridge, Massachusetts.
- Bullard, James**, “Measuring Inflation: The Core is Rotten,” <http://www.stlouisfed.org/newsroom/speeches/pdf/2011-05-18.pdf> 2011. Speech to Money Marketeers of New York University, New York, NY.
- Day, N.E.**, “Estimating the components of a mixture of normal distribution,” *Biometrika*, 1969, 56, 463–474.
- Engemann, K., M. Owyang, and H. Wall**, “Where is an Oil Shock?,” Working Paper 2011-016A, Federal Reserve Bank of St. Louis 2011.
- Gilbert, Christopher L.**, “Speculative Influences on Commodity Futures Prices 2006-2008,” Working Paper 2009.
- Hamilton, James D.**, “A quasi-Bayesian approach to estimating parameters for mixture of normal distribution,” *Journal of Business and Economic Statistics*, 1991, 9, 27–39.
- Kim, Chang Jin**, “Dynamic linear models with Markov-switching,” *Journal of Econometrics*, 1994, 60, 1–22.
- Mankiw, N. Gregory, Ricardo Reis, and Justin Wolfers**, “Disagreement about Inflation Expectations,” Working Paper 9796, NBER 2003.
- Pasaogullari, Mehmet**, “Recent Developments in Inflation Expectations,” <http://www.clevelandfed.org/research/trends/2011/0411/02infpri.cfm> 2011. 01APR2011 entry at Federal Reserve Bank of Cleveland Economic Trends.
- Phillips, Peter C. B. and Jun Yu**, “Dating the Timeline of Financial Bubbles During the Subprime Crisis,” *Quantitative Economics*, forthcoming, 2011.
- Phillips, Peter C.B., Shu-Ping Shi, and Jun Yu**, “Testing for Multiple Bubbles,” *Singapore Management University, Center for Financial Econometrics Working Paper 2010-03*, 2011.
- , **Y. Wu, and Jun Yu**, “Explosive Behavior in the 1990s Nasdaq: When did Exuberance Escalate Asset Values?,” *International Economic Review*, 2011, 52, 201–226.

**Rich, Robert, Joseph Song, and Joseph Tracy**, “Are Rising Commodity Prices Unanchoring Inflation Expectations?,” <http://libertystreeteconomics.newyorkfed.org/2011/05/are-rising-commodity-prices-unanchoring-inflation-expectations.html> 2011. 19May2011 entry at Liberty Street Economics blog.

**Shi, Shu-Ping**, “Bubbles or volatility: A markov-switching ADF test with regime-varying error variance,” *Australian National University, School of Economics Working Paper 2010-524*, 2010.

**Shi, Shuping and Vipin Arora**, “An Application of Models of Speculative Behaviour to Oil Prices,” Working Paper 11/2011, CAMA 2011.

**Trehan, Bharat**, “Household Inflation Expectations and the Price of Oil: Its Deja Vu All Over Again,” *FRBSF Economic Letter*, 2011, 2011-16.

## Appendix

A closer look at the SPF data also shows inflation expectations rising substantially when headline PCE deviates from core PCE explosively in 2007-2008. In addition to point forecasts, each respondent in the SPF survey is asked to assign a probability to different possible values of inflation in the next year. [Rich et al. \(2011\)](#) aggregate these probabilities over respondents, and argue that large changes in the forecasted probabilities of outlying values of inflation provide a signal for inflation expectations. For example, if the respondents’ probability of inflation in excess of 3.0 percent rises substantially from one forecast to the next, there may be cause for concern regarding the anchoring of inflation expectations. That is, changes in the outlying probabilities can act as a leading indicator of possible un-anchoring of expectations ([Rich et al., 2011](#)).

Figure 7

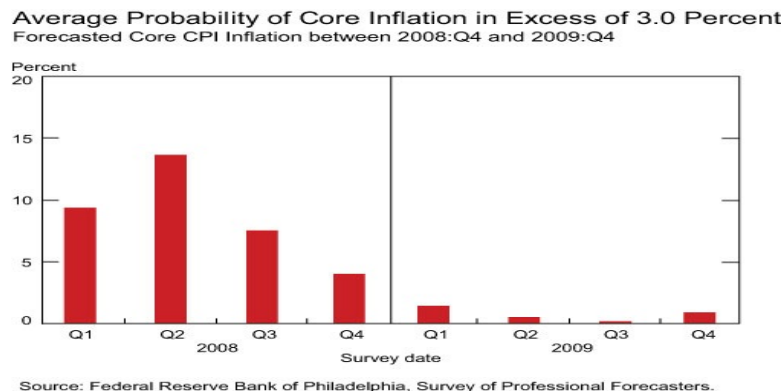


Figure 6 reproduces a plot from [Rich et al. \(2011\)](#). The figure shows the respondents average probability of core inflation in excess of 3.0 percent. It shows a substantial rise in predictions of inflation above three percent during the time that headline PCE was expanding explosively relative to core PCE. This also underlies the importance of the relative rate of change between headline and core measures of PCE.